Is Low Worker Mobility Evidence for Monopsony Power?

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Abstract

In this study, an assessment of the degree of worker mobility in the Portuguese Labor Market is performed by building on the dynamic monopsony literature. By producing firm-level estimates of the elasticity of the labor supply facing the firm, it is possible to conclude that workers have little sensitivity to real wage changes and, thus, that firms enjoy considerable wage-setting power. However, the extend to which one can relate such low worker mobility to monopsony power is severely questioned by first estimating the impact of the estimated measures of the labor supply elasticity on wages and by then assessing how such estimates correlate with three high-dimensional fixed effects - worker, firm and job title - taken from the standard Mincer equation.

1 Introduction

"What happens if an employer cuts the wage it pays its workers by one cent?"

(Allan Manning, *Monopsony in Motion,* 2003)

"If perfectly free competition prevailed everywhere, the wage rate paid by any employer in any occupation would be determinate at a definite point. The value of the marginal net product of labor of given quality would be the same to all employers [...] and, if one employer offered a man less than others, that man would know that he could at once get as much as this value of his marginal net product from others. In so far, however, as movements of workpeople are hampered by ignorance and costs, a monopolistic element is introduced into the wage bargain."

(Arthur Pigou, *The Economics of Welfare,* 1932)

In a perfectly competitive labor market, each worker has full information about the number of jobs available and the wage they pay, is capable of processing such information instantly and faces no costs of switching jobs, leaving competing employers no choice but to offer the individual his/her (known) marginal revenue product. Additionally, firms face no costs from worker turnover. In such a labor market, wage dispersion comes as a consequence of worker and firm heterogeneity, inasmuch as these lead to a set of possible productivity realizations and, additionally, as workers differ in their preferences for non-monetary aspects of their potential employers and as firms differ in the offered combinations of such aspects with the wage they pay. The labor supply facing the firm is perfectly elastic.

By contrast, in an imperfectly competitive labor market wages vary according to many other factors, such as the reservation wages of workers, job offer arrival rates (assumed to be infinite in the competitive model), job destruction rates or the size of the labor market, to name a few. In such a market, there are bad and good jobs and the reasons for this lie beyond productive factors. Considerations regarding the structure of the labor market, such as the number of firms contesting a pool of workers and the degree of inter-firm mobility of such workers, arise as crucial variables of interest in wage determination.

Monopsony was first defined and analytically treated by Joan Robinson (1933), who coined the term so as to replace the formerly used "monopoly-buyer" (Karatzas, 2009). While the actual word literally stands for "single buyer", the concept of monopsony power should be treated the same way monopoly power is unanimously regarded by economists: as the existence of "power over price through control of quantity" (Boal and Ransom, 2002), with the difference that such power comes from the demand rather than from the supply side of the market. Therefore, there is demand market power in the labor market whenever employers have some discretion over the wage they pay or, put in another way, whenever the supply of labor facing the firm is upward-sloping.

This labor supply curve is not to be confounded with the market labor supply curve, as the latter reflects workers' decision on whether or not to participate in the labor market while the former translates their decision, once participating in the labor market, to accept a given job offer or to keep searching while either employed or non-employed.[1] The positive slope of the market labor supply curve arises from heterogeneity in reservation wages - the higher "the market wage", the more hours of leisure workers will be willing to sell or the more workers will want to participate in the labor market -, while the positive slope of the labor supply facing the firm arises because a firm paying a higher wage is more likely to attract workers from other firms. The second concept is, hence, the right one to assess the degree of demand-side competition in the labor market.[2]

Since Robinson, monopsony has been regarded as a mere theoretical issue, with weak adherence to real-life labor markets.[3] However, many explanations of labor market phenomena implicitly assume some kind of wage-setting power by employers, with the special example of search theory; in fact, it was only when economists were able to analytically generate equilibrium wage dispersion that labor market monopsony effectively entered the research agenda. In their seminal paper, Burdett and Mortensen (1998) formulated a non-degenerate equilibrium wage distribution in a wage-posting search model in which workers receive a finite number of job offers per period, i.e., in a model with labor market frictions. This paper, which is considered to be the cornerstone of the new monopsony literature, inspired Manning (2003) to take the model further, using it to provide explanations to many empirical findings in a much more plausible way than those offered by the competitive model. For instance, the gender pay gap can be explained in terms of differences in the elasticity of the labor supply to the firm, with female workers having a less elastic supply due to the fact that they value non-monetary aspects of the job (such as commuting time or number of hours offered) more; in this case, firms will engage in third-degree price discrimination by paying a lower wage to the type of labor with the less elastic supply. Another example regards the returns to experience: in a market where workers have less than perfect mobility, firms will supply general training (as it actually happens) and gains from experience reflect the fact that workers have had more time to work themselves up the job ladder.[4] Yet another theoretical ramification of the assumption that employers post wages,

[1]The clearest example comes from the perfect competition model: a positively-slopped market labor supply coexists with an infinitely elastic labor supply curve facing each firm.

[2]Analogously, it is the firm-level demand elasticity Industrial Organization looks at when addressing the existence of monopoly power.

[3]The little interest labor economists have felt for monopsony is powerfully illustrated by Manning (2003): "The first two volumes of the Handbook of Labor Economics (Ashenfelter and Layard, 1986) contain only two references to monopsony out of a total of 1268 pages [...]. The three subsequent volumes published in 1999 (Ashenfelter and Card, 1999) contain three references in 2362 pages". Manning himself took matters into his own hands, writing the chapter on Imperfect Competition in Labor Markets in the 4th volume of the Handbook.

[4]Returns to tenure, in turn, are explained by stating that more senior workers are those who have already few better options left from which to choose.

while not belonging to the class of monopsonistic models, is efficiency wage theory, which states that firms pay above market-clearing levels so as to avoid shirking by workers or to reduce turnover.

At this stage, a word of notice should be stated: monopsony is not the only theoretical framework where labor market frictions can be found. The statement that the relationship between the employee and the employer creates rents (i.e., that the wage paid is seldom equal to a worker's marginal revenue product) has very predictably generated two general views on the issue, of which monopsony, that focused on the appropriation of such rents by the employer, is one. The other side of the issue views rents being split between the worker and the employer according to the bargaining power of workers, and is based on ex post wage-bargaining models like those exposed in Mortensen and Pissarides (1999). A very competent overview of both sides of the literature on labor market imperfections can be found in Manning (2011).

This study is the second one to produce firm-level estimates of the wage elasticity of the labor supply curve facing the firm (being the first one that of Webber, 2013a), by building on the dynamic monopsony literature introduced in Section 2. Section 3 exposes the theoretical framework sustaining the empirical specification of Section 5, with the dataset being described in Section 4. The results, which will reveal the degree of worker mobility in the Portuguese labor market, will then be fed into a wage equation to assess the impact of such market power upon wages, as well as to study whether there is any relation between them and standard fixed effects. This second analysis of Section 6, as well as the empirical approach to the estimation of the labor supply elasticities, are to our knowledge a novelty in the literature. Section 7 concludes.

2 The dynamic monopsony literature

Since the release of Burdett and Mortensen (1998) and especially after Manning (2003) suggested a way of estimating the average elasticity of the labor supply curve facing the firm (in a steady-state) based on the wage elasticity of separations, there has been some (though not abundant) empirical interest in putting a number to this quantity. This subset of the monopsony literature, which is denoted as that of "dynamic monopsony models", has evolved considerably ever since. Ransom and Sims (2010) and Ransom and Oaxaca (2010) have arguably been the first studies following these lines, although they have not gone any further besides stating the elasticity of labor supply as negative two times the elasticity of separations. Some months after, Booth and Katic (2010) acknowledge that transitions to and from non-employment are also sensitive to wages, and estimate separation equations independently for transitions to/from employment and to/from unemployment.

In the same year, Hirsch et al. (2010) becomes the first study to fully implement Manning's recommendations, recognizing that a firm's separation to non-employment is not another firm's recruit from non-employment. In the meanwhile, Falch (2011, 2013) estimates the average wage elasticities of separations and recruits, respectively, based on a mandated rise in the wages of some demand-constrained schools in Norway. Two large steps further were then taken: first, Depew and Sorensen (2011) manage to estimate the average elasticity of the labor supply to the firm in the short-run; later, Webber (2013a,b) is able to produce firm-specific, short-run estimates, estimating the distribution of monopsony power across American firms. Finally, Hirsch et al. (2013) improves upon Webber's (2013a,b) theoretical model to estimate the cyclicity of labor market power, being arguably the state-of-the-art in dynamic monopsony models. These three papers provide the main inspiration to this study.

3 Theoretical Framework

The empirical strategy which will lead us to firm-specific estimates of the elasticity of the labor supply is anchored to theoretical considerations consisting of a partial equilibrium application of the dynamic monopsony model of Burdett and Mortensen (1998), first presented in Manning (2003) and which have been developed as the dynamic monopsony literature evolved.

We start by assuming that the labor supply faced by the firm is given by the following dynamic identity for employment:[5]

$$N_t = N_{t-1}[1 - s(w_t)] + R(w_t) = N_{t-1}[1 - s^e(w_t) - s^n(w_t)] + R^e(w_t) + R^n(w_t)\,, \tag{1}$$

where N_t is employment at the firm at period t, $s(w_t)$ is the firm's separation rate from period $t-1$ to period t and $R(w_t)$ is the flow of recruits from period $t-1$ to period t. For reasons which will become clear below, we decompose worker flows into flows to/from employment (superscript "e") and flows to/from non-employment (superscript "n"). Letting $\gamma_t = [(N_t - N_{t-1})/N_{t-1}]$ be the periodic employment growth rate and rearranging, one obtains

$$N_t = \frac{[R^e(w_t) + R^n(w_t)](1 + \gamma_t)}{\gamma_t + s^e(w_t) + s^n(w_t)}\,. \tag{2}$$

Now, one takes the natural logarithm on both sides, differentiates and multiplies by the periodic wage to

[5] After which the term "dynamic monopsony" was named. It simply states that employment in one period equals last period's employment times the percentage of workers who do not separate (those who stayed with the firm) plus recruits (those who joined the firm).

obtain the expression in elasticities

$$\varepsilon_{Nw} = \frac{R^e(w_t)\varepsilon^e_{Rw}}{R^e(w_t)+R^n(w_t)} + \frac{R^n(w_t)\varepsilon^n_{Rw}}{R^e(w_t)+R^n(w_t)} - \frac{s^e(w_t)\varepsilon^e_{sw}+s^n(w_t)\varepsilon^n_{sw}}{\gamma_t+s^e(w_t)+s^n(w_t)}$$

$$= \theta_R(w_t)\varepsilon^e_{Rw} + [1-\theta_R(w_t)]\varepsilon^n_{Rw} - \frac{s^e(w_t)\theta_s(w_t)}{\gamma_t\theta_s(w_t)+s^e(w_t)}\varepsilon^e_{sw} - \frac{[1-\theta_s(w_t)]s^n(w_t)}{[1-\theta_s(w_t)]\gamma_t+s_n(w_t)}\varepsilon^n_{sw}\,, \tag{3}$$

where $\theta_R(w_t)$ is the share of recruits from employment and $\theta_s(w_t)$ is the share of separations to employment. At this stage, one has that estimating the elasticity of the labor supply to the firm requires estimation of four worker flow elasticities. While the elasticity of separations can be estimated from a standard matched employer-employee data set like the one used here, to estimate the elasticity of recruitment would require information on the number of applications each firm receives for any given job and at any going wage, which is not available in our data set (nor in any other in the dynamic monopsony literature).

To circumvent this issue, we resort to the Burdett and Mortensen (1998) model so as to relate recruitment from employment to separations to employment. In this model, employed and non-employed workers are offered jobs at a constant rate λ, with wages being drawn from the wage-offer distribution $F(w)$ - the actual wage offer distribution observed in reality, which is the equilibrium variable of the model. Because employed workers are assumed to switch jobs whenever they receive a higher-paying offer, job flows to/from employment when the firm is paying a wage w_t are given by

$$s^e(w_t) = \lambda[1-F(w_t)] \tag{4}$$

$$R^e(w_t) = \lambda\int_{\underline{w}}^{w_t} N(x)dF(x) \tag{5}$$

By (4), the elasticity of separations to employment is given by $[-\lambda w_t f(w_t)/s^e(w_t)]$. Using this expression, (2) and (5), one obtains

$$\varepsilon^e_{Rw} = \frac{\lambda N_t f(w_t)w_t}{R^e(w_t)} = -\varepsilon^e_{sw}\frac{\theta_s(w_t)(1+\gamma_t)s^e(w_t)}{\theta_R(w_t)[\theta_s(w_t)\gamma_t+s^e(w_t)]}\,. \tag{6}$$

To address recruitment from non-employment, Manning (2003) stated that by using the expression for $\theta_R(w_t)$ and applying the transformations required to arrive to recruitment elasticities (taking logs, differentiating and multiplying by w_t), one reaches

$$\varepsilon^n_{Rw} = \varepsilon^e_{Rw} - \frac{w_t\theta^{'}_R(w_t)}{\theta_R(w_t)[1-\theta_R(w_t)]} = \varepsilon^e_{Rw} - \frac{\varepsilon^R_{\theta w}}{[1-\theta_R(w_t)]}\,, \tag{7}$$

where $\varepsilon^R_{\theta w}$ stands for the wage elasticity of the share of recruits from employment. Finally, combining (3), (6) and (7) yields the expression which enables estimation of the elasticity of the labor supply facing the firm:

$$\varepsilon_{Nw} = -\frac{\theta_s(w_t)[1+\gamma_t+\theta_R(w_t)]s^e(w_t)}{\theta_R(w_t)[\theta_s(w_t)\gamma_t+s^e(w_t)]}\varepsilon^e_{sw} - \frac{[1-\theta_s(w_t)]s^n(w_t)}{s^n(w_t)+[1-\theta_s(w_t)]\gamma_t}\varepsilon^n_{sw} - \varepsilon^R_{\theta w} \quad (8)$$

In summary, to estimate ε_{Nw} it is necessary to model the elasticity of separations to employment (ε^e_{sw}), the elasticity of separations to non-employment (ε^n_{sw})[6] and the elasticity of the share of recruits from employment ($\varepsilon^R_{\theta w}$) and to combine them with some computed measures. Equation (8) is different from the one used in Webber (2013a,b) but equal to that of Hirsch et al. (2013), as the former implicitly assumes, upon imposing the theoretical conditions (4) and (5), that $R(w_t)/R^e(w_t) = 1$. One note worth mentioning is that it is not enough to set $\gamma_t = 0$ to obtain long-run estimates of ε_{Nw}, as stated in Webber (2013a) or Hirsch et al. (2013); the three elasticity estimates should come from a model setting one period as the long-run, thereby raising the problems of defining such time horizon and of reduced sample size (through the collapse of the time dimension); we therefore do not compute such measure and restrain ourselves to short-run estimates. The next section now looks at the empirical procedure to estimate the three quantities.

4 Data

In this study, we resort to *Quadros de Pessoal* (Portuguese for "Personnel Records"), an annual administrative linked employer-employee data set started in 1985 (with interruptions in 1990 and 2001). It is a compulsory employment survey covering all establishments with wage earners (aside from independent workers and civil servants) and features, among other data on workers and their employers, detailed information on earnings, namely the base wage but also regular and irregular benefits, overtime pay and the corresponding wage bargaining mechanism. For further information on the data set, please see Cardoso et al. (2012).

The sample for the estimation of firm-level labor supply elasticities spans from 1986 to 2012 and contains all workers aging from 18 to 64 years old, with a wage of at least 80% of the minimum wage and with a single social security code (the worker identifier). Only workers with at least 120 weekly hours worked are included and multiple job holders are excluded. The measure of wage used is real hourly wage, with wages deflated to 1986 currency units according to CPI.[7]

[6]It should be noted that estimating the elasticity of separations to non-employment has as an implicit assumption that there is a stochastic component to workers' reservation wage, in contrast with the simplest equilibrium search models such as Burdett and Mortensen (1998).

[7]By comparison with the previous literature, which uses mainly quarterly earnings, one should note that the time scaling of the wage variable is not relevant *per se*, in that one is measuring coefficients on the percent change of such variable, but

In a slightly more tolerant fashion than that of Webber (2013a), performance of firm-specific estimation is done on a sample comprised of firms with no less than 20 separations to employment, 20 separations to non-employment, 20 recruits from employment and 20 recruits from non-employment over the sample period. This resulted in a sample of 7,331 firms, of which 664 were dropped during estimation of the required elasticities due to lack of convergence in a given model; the remaining firms account for 11,143,220 job spell-year observations and 2,673,869 independent workers, accompanied over the 26-year period. For the estimation of the full-economy models,[8] the last restriction is not necessary and we thus use a sample with 43,209,364 observations (7,061,432 workers). For the estimation of wage equations with high-dimensional fixed effects and for the analysis of the sources of monopsony power, we will resort to the restricted dataset so as to use our firm-level estimates. Finally, specification tests on the models for estimating the three components of (8) were conducted on a sub-sample due to the very demanding computational requirements of estimating three models per each of the 7,331 firms in many different ways; the Aveiro region was thus chosen due to its medium size (988,158 observations, representing 192,764 workers and 494 firms), which makes the exercise feasible and, at the same time, econometrically representative.

Quadros de Pessoal is often regarded as presenting high-quality data, as it is meant for public use and as it is consequently put through regular quality checks by the Ministry of Employment, the public entity responsible for the database. However, there is one variable which is crucial to the estimation of the wage elasticity of labor supply facing the firm and which is measured with error: the indicator variable for whether the worker is a recruit or not. This happens because some workers' first registry under a job spell states a tenure of more than 12 months, which does not qualify such observations as recruits. This has an indirect impact in the estimation of the labor supply elasticities, as the number of recruits at each period is underestimated, which by turn causes net job creation to be underestimated as well. The employment growth rate is hence replaced by the growth rate of a variable reported in *Quadros de Pessoal,* which is the number of people serving the firm (employees and managers) and which correlates very strongly with the computed firm size. There is also a direct effect in estimating $\varepsilon^R_{\theta w}$, though one should note that this does not go beyond diminishing the number of observations for the model which will estimate the wage elasticity of the share of recruits from employment, and should therefore not imply any inconsistency from the empirical estimates.

rather to control for the number of hours worker.

[8] Which are meant to yield the average value of the labor supply elasticity facing the firm.

5 Firm-level labor supply elasticity

Estimation

Section 3 led to the conclusion that three quantities must be estimated for us to measure ε_{Nw}. In what regards the first two, by modeling the instantaneous separation rates to employment and to non-employment as, respectively, $s^E(x) = \exp(\beta^E x)$ and $s^N(x) = \exp(\beta^N x)$, and by assuming that they are, conditional on the vector of covariates x, independent, Manning (2003) has shown that the two may be estimated disjointedly by two univariate models.[9] Bearing this in mind, and following the most recent dynamic monopsony literature, one assumes that the duration of a given job spell (the dependent variable in both models) follows a Proportional Hazards model with exponential scale factor, i.e.,

$$\lambda^h(t,x) = \lambda_0^h(t)\, exp[\beta^h ln(earnings_i(t)) + \varphi^h X_i(t)],\ h = \{n, e\} \tag{9}$$

where λ_0^N is the baseline hazard, t is the length of employment and $earnings_i(t)$ is the individual's base wage, i.e., base wage divided by the number of hours worked during the month.[10] $X_i(t)$ is a set of explanatory variables which includes gender, age, education dummies and year dummies.

While we assume that the data-generating process of our observations is a continuous time one, we acknowledge that the survival time data in our set are grouped - exact survival times are unknown and we only observe the time interval (the year) within which they fall - which leads to the existence of tied durations. Under this scenario, the exact likelihood function becomes quite complicated and both exact and approximate estimation procedures increase computational time considerably. It is possible,[11] nevertheless, to estimate the φ^h and β^h parameters from a discrete time representation of the model: the complementary log-log model; Appendix 1 shows how to go from a continuous-time PH model to this specification. Having said that, our particular models for separations to employment and non-employment are the following:

$$\lambda^h(a_k, x) = 1 - exp[-exp(\beta^h ln(earnings_i(a_k)) + \varphi^h X_i(a_k) + \tau_k^h)],\ h = \{n, e\} \tag{10}$$

where $\lambda^h(a_k, x)$ is the discrete-time hazard function at time a_k, the end of the period of time started at a_{k-1}, and the τ_k describe the pattern of duration dependence. Regarding these parameters, one will test

[9] Equivalently, one could estimate a Competing Risks Model, with employment at another firm and non-employment as the two possible exit routes from a given employment spell.

[10] Total hourly earnings are used in a specification test provided below.

[11] And recommendable, given that the number of failures at each period is far from being of negligible size relative the number of individuals at risk: the economy-wide average separation rate ranges from 14% to 23%.

two different specifications in the Aveiro dataset: a constant baseline hazard, which sets $\tau_k = 0$ for all k, and a nonparametric estimation of the baseline hazard, which sets τ_k as job spell duration-specific dummy variables.[12] It should be highlighted that φ^h and $\beta^h$are the same parameters as those of the continuous time model.

In both models, the coefficient on log earnings is an estimate of the corresponding separation elasticity. During estimation, the entire sample will be used, with workers who remain with the same employer at the end of the data set being considered to have a censored employment spell, as well as job movers for the model in which $h = n$ and those who separate to non-employment when $h = e$. This contrasts with all previous literature, which proceeded like in this study in the model for separations to non-employment but restricted the sample in the model for separations to employment to workers who do not separate to non-employment. While both approaches are valid, one takes a surprisingly original step by behaving coherently in both models.

As for the third quantity, the wage elasticity of the share of recruits from employment, it is estimated through a logit model for the probability that a given new worker was recruited from employment

$$P_{R,emp} = \frac{\exp(\beta^\theta \ln(earnings_{it}) + \varphi^\theta X_{it})}{1 + \exp(\beta^\theta \ln(earnings_{it}) + \varphi^\theta X_{it})} \tag{11}$$

where $P_{R,emp} = 1$ if a recruit comes from employment and X_i contains the same explanatory variables as in the separations models. Regarding our coefficient of interest, one has that $\varepsilon^R_{\theta w} = \beta^\theta * [1 - \theta_R(w_t)]$, with the second factor coming as an observation-weighted time average of the year-firm figures, due to the time-invariant specification imposed upon $\varepsilon^R_{\theta w}$(by turn, for consistency with the other two estimated elasticities) and because years with more observations contribute more to the estimation procedure.

After estimating the three models above, one just needs to compute $\theta_R(w_t)$, $\theta_s(w_t)$, $s^e(w_t)$, $s^n(w_t)$ and γ_t to obtain firm-level estimates of the labor supply elasticity according to equation (8). By using time-invariant coefficients on log hourly real earnings in the computation of the elasticity of the labor supply over the short-run, one implicitly assumes that $s^e(w_t)$, $s^n(w_t)$ and $\theta_R(w_t)$ have iso-elastic specifications and that ε_{Nw} only varies over time according to the weights of ε^e_{sw} and ε^n_{sw}. This is the price to pay for not introducing additional noise in the elasticity estimates, which would come as a consequence of replacing a single regressor with T_j interaction terms, where T_j is the number of sample periods during which one observes firm j.

[12]While the pattern of duration dependence in the underlying hazard function cannot be described without imposing restrictions on the specification of the τ^h_k, these will represent the change in the log of the integrated hazard function from the end of one interval to the other - i.e., the regression coefficients will be estimated in a semi-parametric fashion, in an almost analogous way to the estimation of a Cox PH model in continuous time.

The results from estimating several specifications of (10) and (11) for the Aveiro dataset are presented in the following table:

	ε_{sw}^{e}	ε_{sw}^{n}	$\varepsilon_{\theta w}^{R}$	ε_{Nw}	$\varepsilon_{Nw}(full)$
Constant baseline hazard					
Wage only	-.4322	-.6563	.5553	.5394	1.2018
No year effects	-.5509	-.8436	.7612	.6316	1.3363
Full model	-.7016	-.7218	.9001	.5705	.5217
Full model, total earnings	-.3910	-.5030	.6969	.2155	.4443
Nonparametric baseline hazard					
Wage only	-.0023	-.4388	.5553	-.2044	-.4426
No year effects	-.1914	-.5606	.7612	-.0712	-.6643
Full model	-.2789	-.4401	.9001	-.3592	-.1020
Full model, total earnings	-.0588	-.2608	.6969	-.3708	-.2152

Estimates for the Aveiro dataset. The 2.5% largest and smallest estimates of ε_{Nw} were removed from the sample. Firm-level averages are presented in the first four columns, time-average values from full-sample estimation presented in the fifth column.

Table 1: Firm-level average of estimated elasticities of the labor supply.

The results displayed in Table 1 present a sizable difference between the models assuming a constant baseline hazard and those which rely on nonparametric estimation of this component of the hazard function (and which show that separations to both employment and non-employment display negative duration dependence). The effect of allowing for duration dependence on the estimated elasticities is in line with what is predicted by Manning (2003, p.103): "The inclusion of job tenure always drastically reduces the estimated wage elasticity as high-tenure workers are less likely to leave the firm and are more likely to have high wages"; this is due to the fact that, for a higher wage, it is less likely that workers will get a higher offer elsewhere and it is less likely for the reservation wage to rise above the wage. As we believe that controlling for tenure will capture the effect of wages on the likelihood that workers will separate - meaning, in the light of our theoretical framework, the likelihood that a worker gets a higher wage offer or the likelihood that his/her reservation wage rises above its wage -, which is precisely the effect of interest, we will use the estimates from the exponential model.

Moreover, and perhaps most importantly, it should be highlighted that our choice of a duration model is driven by its functional match with Manning's (2003) exposition of the separation rates as instantaneous (thus, exponential) rates, rather than by a belief that this is how the issue should be genuinely modeled. What we want to obtain from the separation equations is the wage elasticity of the separation rate which, by the time one considers controlling for tenure, becomes a different empirical question from the wage elasticity of the (discrete) hazard function for a given worker. This means that, on sheer econometric grounds, estimating the wage elasticity of the separation rate through any other binary model, such as a logit or a probit model, would be equally valid. The advantage of using the complementary log-log model is that one can map our procedure directly to the specific empirical issue at stake - our coefficients of interest are the desired elasticities - and to the existing literature.

Regarding other specification issues, one may state a big contrast relative to Webber's (2013a) low sensitivity to the inclusion of controls to the models. Both standard individual controls and year effects (capturing the impact of the macroeconomic environment in the firm's separation rate) lead the coefficients on the wage to vary significantly, both in the constant and in the nonparametric hazard specifications. Such sensitivity is also an issue when regarding the choice between using the base wage or total wage as the independent variable of interest in the three models. One can argue in favor of each of these variables: on the one hand, insofar as total wages include variable payments, which appear in our database as absolute payments, using them as the wage variable will introduce some wage variation to which the worker will never react (that which is due to volatility in worker performance) and will, hence, underestimate the labor supply elasticity; on the other hand, using the base wage as the wage variable will ignore worker heterogeneity in terms of their compensation schema. We choose in favor of the one whose imperfections have an ambiguous, rather than decisively downward-biasing, impact, which is also the most conservative measure, i.e., that which leads to a higher estimate for the elasticity - the base wage.

Results

After obtaining our firm-level measures, we erased from the sample the top and bottom 1% of the estimates. Moreover, we made the choice of excluding from the sample firms whose time-average of the estimates for ε_{Nw} came out negative. It is our understanding that the fact that we cannot distinguish quits from fires in our dataset can jeopardize our purposes of identifying the labor supply, in that it rather captures labor demand determinants. Under the innocuous theoretical prediction that, everything else constant, workers with a higher wage are more likely to be laid-off by employers (as it is more likely that a negative shock drives match productivity below the wage, the main driver of job destruction in search models like those of Mortensen and Pissarides, 1994), firms whose corresponding separations disguise a high share of quits will present positive separation elasticities, which by turn will contribute to a smaller estimate for the elasticity of the labor supply. Therefore, our decision aims at keeping only those firms whose separations represent, with a good level of reliability, voluntary quits rather than dismissals.

Having said that, we proceed by presenting our estimates for ε_N in the following histogram, with summary statistics presented in Table 2 below:

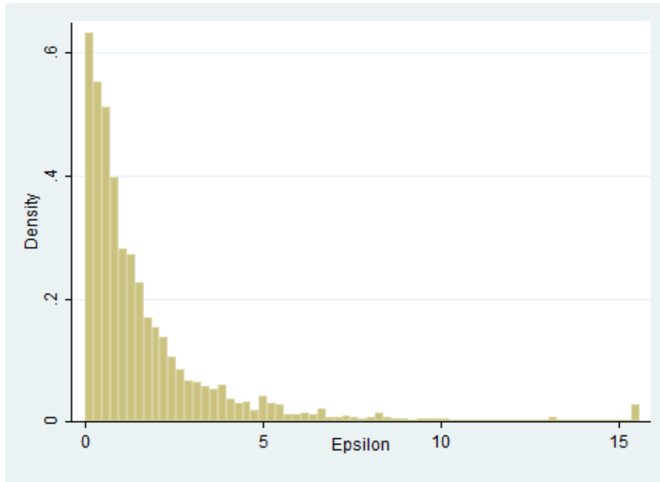

Figure 1: Histogram of the estimated wage elasticities of the labor supply facing the firm.

	1%	5%	10%	25%	50%	75%	90%	95%	Mean	Full
ε_N	0.018	0.085	0.161	0.415	0.977	2.130	4.269	6.446	1.804	1.502

Figures for the observation-weighted distribution of estimated elasticities. Time-average figure for the full model.
Firm-level estimates: 6,017,923 observations | Full economy model: 43,209,364 observations

Table 2: Distributional statistics for the estimated wage elasticities of the labor supply facing the firm.

This said, one should note that the firm-level average figure of 1.804 - as well as the full economy figure, 1.502 - reveals that firms possess a large degree of monopsony power in the Portuguese labor market. These estimates should be compared to those of Webber (2013a) and Hirsch et al. (2013), which deploy the same empirical methodology to countrywide matched employer-employee data sets. We find that our full economy estimate is much lower than Hirsch et al.'s (2013) 2.044[13] and higher than Webber's (2013a) 0.76, while the average value for the firm-level estimates also ranks above that of the latter paper, 1.08. Comparisons with this study are, however, deemed to be inconclusive, as the author's empirical approach to the estimation of the separation elasticities (Cox Proportional Hazards model) controls for tenure, which the effects highlighted above. One can only state that workers in Portugal are, on average, less mobile than their German counterparts, something that fits well in the common understanding one has of contemporary labor markets in advanced economies (Elsby et al., 2013), and that our firm-level estimates are, as Webber's, concentrated in very low values, with values close to zero as the mode. Perhaps most importantly, another crucial difference between this work and our two main references is that our data are annual (rather than quarterly), which attenuates the observable frequency of job and worker flows and also biases our results downward in the presence of unobserved worker heterogeneity, in light of what is

[13]Average value of the time-varying estimates, where the three elasticities are themselves time-varying as the wage was interacted with the lagged unemployment rate.

shown in Manning (2003, p.110).

One can further make use of our firm-level estimates to see how they vary along sectors. One obtains this by running a regression of the estimates on a set of REV2 sector dummies with no constant to obtain the average value of the firm-level estimates per sector. In Table 3 below, one can see that more unionized sectors appear to be the ones with the most elastic labor supply (Transportation and Communication, Public Enterprises or Financial Services), while sectors such as Retail or Chemical Products are where one can find the least mobile workers.

Sector	ε_N
Fishing	2.913
Energy Products	2.327
Mining	1.853
Food, Beverages and Tobacco	2.016
Textiles	1.747
Chemical Products	1.225
Metal Industry	1.884
Manufacturing	1.987
Construction	1.962
Retail	1.575
Transportation and Communication	2.083
Financial Services	2.282
Real Estate	1.551
Education	1.651
Health	1.672
Public Administration	2.932

Table 3: Regional average of the wage elasticities of the labor supply facing the firm.

Now that we have analyzed our results from the firm-level estimation exercise, we will submit our results to a specification test by the form of estimating the impact of firm demand market power on workers' wages.

6 Monopsony and wages

Having produced firm-level estimates of the elasticity of labor supply, one can use such estimates to study their impact on wages, following Webber (2013a), and to assess the extend to which fixed effects arising from a wage equation reflect such market power. This second level of analysis is a novelty in the literature and may reveal to what extend different dimensions of unobserved time-invariant heterogeneity correlate with the estimated quantities.

The general model to be estimated in this section is the following:

$$ln(earnings_{ijt}) = \psi\varepsilon_{jt} + \varsigma_t + \delta Z_{it} + \mu F_{jt} + u_{it} \tag{12}$$

where ε_{jt} stands for the estimated labor supply elasticities, ς_t are time effects, Z_{it} is a set of individual controls including tenure and tenure squared, age and age squared, gender and education dummies, F_{jt} is a set of firm-level controls comprised of log real sales per worker, industry and region dummies and log firm size and - finally - u_{it} is the standard error term.

The exercise of market power

In this subsection, interest lies solely upon ψ, the impact of firm market power on wages, and on the coefficient on log size, commonly known as the Employer Size Wage Effect (ESWE). Table 4 presents estimates of the two coefficients for various specifications. All models present in this table feature only year fixed effects, which are suitable to be estimated through the introduction of dummy variables, as the estimated components of ε_{Nw} are time-invariant and, hence, the introduction of firm or worker fixed effects would capture their impact on wages and would render the elasticity of labor supply statistically insignificant for wage determination. Whenever the estimated elasticities are used in estimation, bootstrapped standard errors are computed to account for measurement error in ε_{jt} for evaluating the precision of $\widehat{\psi}$.

	No tenure	No logsize	No market power	Full Model
ε_{jt}	-.0032*	-.0040*		-.0039*
	(.000060)	(.000064)		(.001357)
$ln(N_j)$	.0121*		.0098**	.0013*
	(.000489)		(.004544)	(.005723)
Observations	6,014,491	5,984,026	11,041,801	5,984,026

Standard-errors in parenthesis; Bootstrapping is used whenever ε_{jt} is included; *p<0.01; **p<0.05; ***p<0,1.

Table 4: The impact of market power on wages

Looking at the presented results, the main conclusion is that, regardless of the specification, the impact of the elasticity of the labor supply on earnings is of negligible size - at its largest estimate, an increase of one unit in the elasticity of the labor supply to the firm will lead to a decrease in the wage of approximately .4%. Although economically insignificant in its size, that the coefficient is statistically different from zero on the negative, rather than on the positive size, casts doubt on the validity of our estimates as measuring monopsony power. As explained in Webber (2013b), a negative coefficient estimate reflects an efficiency-wage perspective of the labor market according to which firms would pay workers higher wages to reduce turnover. In such a case, the causality relationship is inverted: firms pay wages above market-clearing levels in order to have lower turnover and to make workers less sensitive to further wage changes. Being it as it may, this estimate comes at a stark contrast to that same study's estimates, which range from 0.13 to 0.15.

Another interesting conclusion one can draw from these results concerns the way the Employer Size-Wage Effect coexists with the estimated elasticities in this wage regression. Although both coefficients are reduced in size when the other is introduced in the equation, both remain statistically significant at the most solid level. This may point out to a labor supply whose components are not iso-elastic or to the fact that the ESWE contains factors other than an upward slopping labor supply curve, such as rent-sharing, compensating wage differentials or unobserved labor quality[14]. Taking Manning's (2003) exposition in consideration, the ESWE may also contain factors affecting labor demand, such as economies of scale or the correlation between labor demand and supply shocks. What seems to be certain is that other factors are relevant in explaining the positive correlation between wages and firm size and that their effect upon wages is much more important in wage determination.

Monopsony and Unobserved Heterogeneity

In this subsection, our main goal is to assess to what extend do the labor supply elasticity estimates are correlated with the standard fixed effects found in wage equations. To this end, we estimate a Mincer equation with no measure of labor market power so as to recover standard fixed effects, against which our measures of market power will be compared.

The model we will estimate is a wage equation with three high-dimensional fixed effects: worker fixed effects (ω_i), firm fixed effects (η_j) and job-match fixed effects (ρ_s), relying on the identification procedure proposed by Guimarães and Portugal (2010), an iterative procedure yielding exact OLS estimation of the high-dimensional fixed effects.

Specifying an index s for the occupational type of the worker, the model is specified as follows. All the control variables remain in such equations (except, of course, those whose identification is made impossible by the inclusion of fixed effects).

$$ln(earnings_{ijt}) = \omega_i + \eta_j + \rho_s + \varsigma_t + \delta Z_{its} + \mu F_{jt} + u_{it} \tag{13}$$

[14] Manning (2003) lists these three factors as possible alternative explanations to the ESWE, stating, respectively, that more productive firms tend to be larger, that larger firms may have worse working conditions or that larger firms have more productive matches.

After estimating the model and recovering the estimates for the three fixed effects, one will assess the correlation between the estimated elasticities and the three layers of unobserved heterogeneity we specify: worker time-invariant characteristics which impact the wage it earns, commonly taken as its ability (unobserved to the analyst but perceived by the employer); firm effects capturing its wage policy and other permanent aspects, such as wage-enhancing technology; job title effect, with job title being the combination of the professional category of the worker with the collective agreement covering his contract, and which reflects "the distinct tasks performed by workers that define the set of occupational boundaries" (Torres et al. 2013).

The results of running an OLS regression of the estimates for the firm-level elasticities on the three fixed effects (plus a set of industry , regional and year controls) are presented in Table 5. Bootstrapped standard-errors are used so as to account for measurement error in the fixed effects.

	No controls	Industry, Region and Year Controls
$\widehat{\omega}_i$	.0385*	.0966*
	(.002872)	(.003141)
$\widehat{\eta}_j$	-.0480*	-.5075*
	(.006110)	(.007076)
$\widehat{\rho}_s$	-.1021*	-.7261*
	(.002795)	(.007136)
Observations	5,984,026	5,984,026

Bootstrapped s.e. in parenthesis. *p<0.01; **p<0.05; ***p<0,1.

Table 5: OLS regression output of ε_{Nw} on firm, worker and job title FE.

The estimates presented in Table 5 point towards the direction of efficiency wages suggested in the last subsection. In both specifications, it is seen that firms with a more generous wage policy face a less elastic labor supply and that job titles which pay more than average can be found in those same firms. It is important to mention that the job title fixed effect is a mix of two effects: the bargaining power of the agents who set minimum wages while discussing each collective agreement and the effect of promotions from one professional category to another. Should this second effect be the most prevalent, one can interpret the correlation between this fixed effect and the the estimated elasticities as extra evidence of the efficiency wage perspective, in that workers experience higher tenure because they are paid more and that leads to promotions within the firm. By contrast with these two conclusions, it seems that more able workers are allocated to firms which face more elastic labor supplies, which may indicate that such workers face less mobility costs and therefore contribute to higher estimates of ε_{Nw}. Insofar as worker fixed effect is the best proxy one can obtain for worker (and, thus, job match) productivity, this result is also evidence against the theoretical model of Postel-Vinay and Robin (2002), which predicts that most productive workers, who will match the most productive firms, would be allocated to the firms with largest wage-setting power. Finally, one should observe that none of the regressors varies with time, by contrast with the regressand, thus the coefficient estimates will reflect the correlation between the fixed effects and some observation-weighted time average of the firm-level elasticities.[15]

7 Conclusion

In this study, we built on the dynamic monopsony literature so as to estimate the degree of worker mobility characterizing the Portuguese labor market by deploying, for the first time, a discrete time specification. After providing evidence on the differences between imposing no duration dependence and introducing tenure non-parametrically in estimating the two separation elasticities, we concluded that employers face very rigid labor supply curves, translating into a high degree of wage-setting power.

Following this exercise, however, we found evidence against the monopsonistic view of the labor market, namely in its prediction of the correlation between the elasticity of the labor supply and wages: we found evidence of a tiny but significant negative correlation between the two variables. While one can argue that it does not mean much to achieve statistical significance with 6 to 11 million observations, the fact is that the data rejected a decisively positive impact of the wage elasticity of the labor supply facing the firm on wages. Moreover, estimating worker, firm and job title fixed effects and analyzing the allocation of such effects along the estimated elasticities reinforced the suggestion of the estimation of the wage equation towards an efficiency wage explanation of the labor market: firms pay workers higher wages to keep them from separating, leading to a negative correlation between the firm's wage policy and the elasticity of the labor supply it faces. Evidence from this is found in our exercise.

In this sense, one may say that while this finding is of utmost importance for the understanding of the functioning of our labor markets, it serves positive economics much better than normative economics.This would be the case even if we would find evidence in favor of the exercise of market power by employers: given what is discussed in Manning (2003), the simple recipe of imposing the right minimum wage upon an isolated monopsonist does not extend to more realistic models of labor markets with frictions. For instance, according to the problem one wants to study in an extended Burdett and Mortensen model, one can have that the free market has too much or too little employment, and there can even be no theoretically optimal minimum wage a policy-maker can set; one just has to introduce simultaneously heterogeneity in the reservation wage of workers and entry (with costs) by firms to have a welfare function with ambiguous

[15] In fact, we also estimated the model using the time average of ε_{Nw} as the dependent variable, obtaining qualitatively similar results.

dependence upon employment - in a market equilibrium with too many firms paying entry costs, workers may be benefited due to the increased wage. Alternatively, introducing human capital investment by workers and firms will generate multiple equilibria. If one adds to this our conclusion regarding the (lack of) exercise of market power by firms upon wages, claiming for labor market intervention on the grounds of the existence of considerable monopsony power becomes an unsurpassable task.

Despite the fact that we were not able to prove monopsony, its main premise is common with what we find: workers are not infinitely mobile across firms, which enables firms to have some discretion over the wage they pay. Going back to the initial words of this study, the answer to the question "What happens if the employer raises the wage it pays its workers by one cent?" is not "It would meet an infinite supply of labor". In that the importance of a labor market with frictions as described by the monopsonistic view does not depend on the existence of a downward pressure, by the employer, on its workers wages, our results do not pinch the potential gains of taking a monopsonistic view to labor markets. In essence, what hinders behind our final conclusions is the same that would be present should we reach conclusions in line with those in Webber (2013a): that workers are not perfectly mobile across potential employers. One should not forget that the lack of worker mobility is precisely what we first expect to - and do - estimate.

All in all, this study has taken a couple of steps further in understanding the nature and sources of demand market power in contemporary labor markets. In our understanding, further research is needed in improving the estimates for the labor supply elasticities, namely by adopting a new theoretical framework based on Burdett and Mortensen (1998) but which may handle at least one of the caveats of the "incumbent" model. One of such flaws concerns the specification of the dynamic labor supply in equation (1); Falch (2013) points out to the fact that this is, in fact, an employment identity which ignores all workers which would like to work for the firm at the given wage but are unable to do so because employment is demand constrained. For firms at which this is the prevailing case, our specification of the labor supply is actually right censored at the employment level, which will lead to an underestimate of the labor supply elasticity. A second caveat concerns the inclusion of instruments other than the wage that the firm may use to impact the labor supply it faces. Manning (2006), in his *Generalized Model of Monopsony*, presents the concept of the labor cost fuction (non-wage per-worker costs of maintaining employment at a given level, such as hiring effort) to reach the effective labor supply elasticity including such factors. This may be the solution for the fact that mandated wage rises and mandated employment increases have both produced small changes in the other endogenous variable across the empirical literature (Manning, 2011). A third problem of our model is the assumption of the Burdett and Mortensen (1998) model that a worker will leave his/her firm for the tiniest wage gain, although this is not considered to be neither crucial nor of straightforward solution, at least in a way that would bring much improvements upon the current state of things. The first two avenues appear as the most promising in providing estimation of labor supply elasticities of the highest quality and, hence, in answering to the big question mark our results, when compared with those of Webber (2013a), have left about demand market power in the labor market.

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Appendix 1: The complementary log-log model

Following Jenkins (2004), synthesizing the combination of covariates and parameters as x and suppressing, for now, the superscript h, the survivor function at time a_k, the end of the interval $(a_{k-1}, a_k]$, is given by

$$S(a_k, x) = exp[-\int_0^{a_k} \lambda_0(u)\phi(x)du] \tag{14}$$

which, setting $\phi(x) = exp(x)$, yields

$$S(a_k, x) = exp[-exp(x)\int_0^{a_k} \lambda_0(u)du] = exp[-H_k exp(x)] \tag{15}$$

with $-H_k \equiv H(a_k) = \int_0^{a_k} \lambda_0(u)du$ being the integrated baseline hazard at the end of the given interval. Looking now at the discrete hazard rate $\lambda(a_k, x) \equiv \lambda_k(x)$:

$$\lambda_k(x) = \frac{S(a_{k-1}, x) - S(a_k, x)}{S(a_{k-1}, x)} = 1 - \frac{exp[-H_k exp(x)]}{exp[-H_{k-1} exp(x)]}$$

$$= 1 - exp[exp(x)(H_{k-1} - H_k)] \tag{16}$$

Finally, applying a $ln(-ln(.))$ transformation leads to the complementary log-log model

$$ln[1 - \lambda_k(x)] = exp(x)(H_{k-1} - H_k)$$

$$\iff ln(-ln[1 - \lambda_k(x)]) = x + ln(H_k - H_{k-1}) = x + \tau_k \tag{17}$$

By estimating a cloglog model, one is able to estimate, from grouped data, the parameters contained in x of the underlying continuous time model, plus the parameters τ_k representing the differences in values of the integrated baseline hazard from the start to the end of the time interval. As stated in the text, while theτ_k cannot describe the exact pattern of duration dependence in the underlying model unless they are assumed to have a given functional form - rather, they identify the pattern of duration dependence in the interval hazard function -, there are cases, like ours, when such pattern is not the object of the study and when interest lies solely upon the estimation of the regression coefficients.